HEALTH INSURANCE AND JOB MOBILITY: 
THE EFFECTS OF PUBLIC POLICY ON JOB-LOCK 

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The authors study a policy of limited insurance portability that has been adopted by a number of states and the federal government over the past 20 years. They find that these "continuation of coverage" mandates, which grant individuals the right to continue purchasing health insurance through their former employers for a specified period after leaving their jobs, are associated with a significant increase in the job mobility of prime age male workers. This finding suggests that "job-lock"—lack of mobility out of jobs that offer health insurance—arises in large part from short-run concerns over portability rather than from long-run problems.

The predominant feature of private health insurance in the United States today is its link with employment. Workplace pooling of risks has several advantages—it allows for economies of scale in administration, reduces the problem of adverse selection as long as workplaces are chosen independently of health status, and allows individuals to take advantage of the non-taxation of employee fringe benefits—but it also has an important disadvantage: if wages do not perfectly offset differences in the valuation of health insurance across different jobs, individuals may not change jobs even when new employment opportunities with higher match-specific productivity arise—a phenomenon known as "job-lock." Recent research has demonstrated the empirical significance of this phenomenon. For example, Madrian (1994) estimated insurance-induced reductions in mobility of approximately 25%. Given a distortion of this magnitude, it is obviously important to ask how public policy can alleviate job-lock.

Three possible strategies to alleviate or eliminate job-lock are to remove the link between health insurance coverage and the employment relationship (by, for example, establishing universal coverage); maintain employment-based insurance but ensure full portability of coverage across jobs; or continue the provision of coverage through employers, but forbid the exclusion from coverage of pre-existing medical conditions and conditions discovered in a mandatory initial physical examination (two practices

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EFFECTS OF PUBLIC POLICY ON JOB-LOCK

that have become increasingly widespread). We argue, however, that implementing any of these relatively comprehensive solutions may be complex and costly.

In this study, we examine instead a partial corrective that has already been used for two decades: "continuation of coverage" mandates, which require employers to continue providing health insurance coverage to workers who leave the firm for a specified period of time. In recent research, we have found that these mandates have large effects on the retirement behavior of older workers (Gruber and Madrian 1993a, 1993b). Our goal in this paper is to determine whether they have also been effective in alleviating job-lock. If they have, they may offer a way to reduce job-lock and still retain the advantages of employer-provided health insurance coverage.

We are able to study the effects of continuation of coverage mandates because of the substantial variation in their timing and generosity. Continuation of coverage laws were passed at different times and in different states from the mid-1970s through the mid-1980s before being federally mandated in 1986. In addition, some laws were quite liberal, allowing 15–20 months of coverage, whereas under others the extent of coverage was minimal. We use data from the Survey of Income and Program Participation (SIPP) to study the effects of these laws on job mobility and subsequent insurance coverage.

Health Insurance and Job Mobility

The primary problem with workplace pooling of risks is that a workplace is not necessarily chosen by workers independently of their health status. This fact has led many insurers to exclude pre-existing conditions from coverage for a period after an individual is hired, typically between six months and two years.1 The New York benefits consulting firm Foster Higgins reports that 57% of employers excluded pre-existing conditions in their health plans in 1987 (Cotton 1991). Furthermore, almost half of full-time workers in firms of 250 employees or more face a length of service requirement before being eligible for any health insurance coverage (Bureau of Labor Statistics 1989). Although the length of service requirement is generally short, it may be important for the employee anticipating major medical expenses.2

In addition, for small firms with experience-rated insurance premiums, hiring a sick employee can entail large increases in insurance costs. As a result, employees with medical problems may face discrimination when they seek a new job with coverage. Even if there is not explicit discrimination in hiring, many insurance policies at small firms are now medically underwritten so that sick employees may not be able to qualify for coverage. Furthermore, only 60% of firms with fewer than 100 employees even offer insurance, so a new job with insurance coverage is far from guaranteed for the potential job changer (Gruber, forthcoming).

The theory of compensating differentials predicts that in a perfectly competitive labor market, differences in the generosity of benefits across firms will be offset by differing wage levels or other aspects of compensation. As Rosen (1986) noted, however, the compensating differential for a given health insurance package will equate the wage reduction to the valuation of that

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1A pre-existing condition is generally defined as any medical problem that has been treated or diagnosed within the past six months to two years. In some cases it may be more broadly defined as any medical problem for which an individual has ever received care. It may also be extended to include medical conditions

2Pre-existing conditions exclusions and length-of-service requirements are likely to pose a greater deterrent to job change as individuals get older and the incidence of chronic medical conditions increases. Our analysis controls for age effects, however, and we consider only those age 54 and younger. See Gruber and Madrian (1993a, 1993b) for a discussion of the health problems of older workers and the effect of continuation coverage on retirement.
package by the marginal worker who is just indifferent between wages and benefits. In a heterogeneous population, some workers will value health insurance more highly than the marginal worker; these “infra-marginal” workers will receive a surplus from their excess valuation of the benefit. That is, these workers will value total compensation at their current job at above the sum of their wages plus the cost to the employer of providing the benefit. As a result, these workers may turn down alternative employment opportunities in which their wages and total productivity would be higher but their health insurance inferior to that in the current job.

The potential market failure in this case arises from the lack of a complete market for worker/job-specific compensation packages. If firms could accurately determine worker valuation of benefits, and if they could offer each individual a customized wage and benefits package, then workers would move to jobs in which their productivity was highest.

Offering worker-specific compensation packages is generally not feasible, however, for at least three reasons. First, such packages would violate a number of rules designed to guard against discrimination in the design of benefit plans. The Internal Revenue Code gives favorable tax treatment to employer expenditures on health insurance only if all workers are provided with an equivalent benefits package. Second, even if worker-specific compensation were allowed, problems with eliciting true worker valuation of these benefits would make it difficult to design. Finally, the costs involved in administering such packages might be large enough to absorb most of the rents that the worker would earn from their existence.

Although this market failure will exist for any job amenity, such as workplace safety, pensions, or parking, there are two reasons why it may be more important for health insurance than for other job amenities. First, the variance in employees’ valuation of health insurance is, in general, much greater than the variance in their valuation of other benefits, due to both the high cost of medical care and the substantial variation across individuals in health status. Individual medical expenditures among those aged 20–54 averaged $1200 in 1990, with a standard deviation of $3878. Large differences in the productivity of an employee across different firms may therefore not be sufficient to induce mobility if other firms do not provide health insurance or if the insurance excludes pre-existing conditions or is medically underwritten.

Second, there are very large differences across firms in the cost of providing health insurance, particularly by firm size. The cost of insurance is estimated to be 35% lower in the largest (10,000+ employee) firms than in smaller (1–4 employee) firms (Congressional Research Service 1988). This difference will make it difficult for small firms to offer insurance, even if they know that they could attract much more productive workers by doing so.

The welfare consequences of health insurance-induced job-lock are unclear (Madrian 1994). Many studies have found large increases in wages for workers who switch to new jobs (Minter 1992; Bartel and Borjas 1977; Mincer 1986; Topel 1986). These higher wages may represent a gain to society from the increased productivity of an improved job match. On the other hand, it may be that some jobs offer rents to workers, and individuals tend to move into these jobs as their careers advance. Furthermore, a key consideration is the temporal nature of the job-lock. Job-lock arising from chronic illnesses could have major and long-lasting consequences for work-
ers’ job productivity, whereas job-lock linked to short-term conditions such as pregnancy may be less important from a life-cycle perspective.

Finally, it can be argued that job-lock allows firms to reap the returns from general human capital investments by reducing the ability of workers to use those investments elsewhere. However, job-lock seems a particularly inefficient mechanism for achieving this goal, relative to back-loaded compensation devices such as pensions, since presumably it is not just the sickest workers that the firm wishes to retain.

Possible Solutions to Job-Lock

A solution to job-lock that suggests itself on first consideration of the problem is to divorce health insurance coverage from the employment relationship. Pooling could occur, instead, along other dimensions, such as regionally (as proposed in Diamond 1992) or nationally (as with Canadian National Health Insurance). As underscored by the current debate about alternative national health care plans, however, policies that change the fundamental nature of health insurance coverage in the United States, where 83% of all private insurance is provided through the workplace (Gruber, forthcoming), have enormous transition costs that may make them infeasible.

Another alternative is to maintain employment-based insurance but to ensure portability of coverage across jobs. A policy of full insurance portability, however, raises a number of tricky design issues. For example, will employees who change jobs frequently be covered by their first employer, each employer in succession, or the employer of their choice? Each of these policy options has difficulties. Making the initial employer responsible for all future insurance would lead to large distortions in initial hiring decisions and massive administrative costs in tracking employees through their job changes. Assigning responsibility to each subsequent employer would make it difficult to guarantee individuals fully comparable coverage on each job, and if coverage is less generous in alternative employment, the problem of job-lock remains.

Perhaps the least disruptive option would be to continue having individuals receive insurance from their current employer but to forbid both the exclusion from coverage of pre-existing conditions and the growing practice of medical underwriting. Those prohibitions would assure individuals of coverage when they change jobs. Unfortunately, they might also make it infeasible for many small firms to offer insurance, since one very sick employee with sufficiently high medical expenses could drive the firm out of business. Thus, with no other government action, many small firms might cease to offer health insurance at all. Even if all firms were required to provide coverage, mechanisms would be needed to reinsure small firms against these major shocks to their risk pools. Thus, effective implementation of this policy option may require much more than simple insurance market regulation.

The key point is that the design of public policies to alleviate labor market distortions without radical change to our health care system is an important and difficult issue. In considering various alternatives, it is useful to reflect on our experience with the major policy intervention to date that addresses this problem: the limited portability of health insurance provided by "continuation of coverage" mandates.

Alleviating Job-Lock Through Continuation Benefits

State and federal continuation of coverage laws require that employers sponsoring group health insurance plans offer terminating employees and their families the right to continue their health insurance coverage through the employer’s plan for a specified period of time. Although individuals must pay the full average cost of their group insurance, the price may be well below that of a policy purchased in the individual market, especially for individuals with high medical expenditures. Thus,
continuation benefits provide a way for workers who leave their jobs to maintain their health insurance coverage for at least a few months, when the alternative would most likely be to go uninsured.

Continuation of coverage laws generally apply to all separations (except those due to an employee's gross misconduct), although in some states benefits are restricted to those who leave their jobs involuntarily. They often also provide benefits to divorced or widowed spouses and their families. The first such law was implemented by Minnesota in 1974. More than 20 states passed similar laws over the next decade before the federal government, as part of its 1985 Consolidated Omnibus Budget Reconciliation Act (COBRA), mandated such coverage at the national level.

The various state statutes are summarized in Table 1. The length of coverage is generally quite short, from three to six months, although ten states mandate coverage of nine months or more. Most state laws stipulate that an employee must have been covered by an employer's insurance for three to six months before being eligible for continuation coverage. The state laws also apply only to firms that actually purchase insurance through an insurance company; self-insured firms, under the 1974 Employee Retirement Income and Security Act (ERISA), are not subject to these (or any other) state mandates.

Although similar in spirit, the state and federal laws differ in a number of important ways. First, the length of coverage mandated under the federal law, 18 months, equals or exceeds that mandated by all but one state (as of January 1987, Connecticut law has provided for up to 20 months of coverage).

Second, there is no minimal length of time for which an employee must be covered under an employer's plan before being eligible for continuation benefits. Third, the federal law applies to self-insured firms, which are exempt from the state laws, as well as to those who purchase their coverage from insurers. The federal law, however, does not apply to small firms employing fewer than 20 workers. Finally,

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### Table 1. State Continuation of Coverage Laws.

<table>
<thead>
<tr>
<th>State</th>
<th>Effective Date</th>
<th>Months of Coverage</th>
</tr>
</thead>
<tbody>
<tr>
<td>Arkansas</td>
<td>7/20/79</td>
<td>4</td>
</tr>
<tr>
<td>California</td>
<td>1/1/85</td>
<td>3</td>
</tr>
<tr>
<td>Colorado</td>
<td>7/1/86</td>
<td>3</td>
</tr>
<tr>
<td>Connecticut</td>
<td>10/1/75, 1/1/87</td>
<td>10, 20</td>
</tr>
<tr>
<td>Georgia</td>
<td>7/1/86</td>
<td>3</td>
</tr>
<tr>
<td>Illinois</td>
<td>1/1/84, 8/25/85</td>
<td>6, 9</td>
</tr>
<tr>
<td>Iowa</td>
<td>7/1/87</td>
<td>9</td>
</tr>
<tr>
<td>Kansas</td>
<td>1/1/78</td>
<td>6</td>
</tr>
<tr>
<td>Kentucky</td>
<td>7/15/80</td>
<td>9</td>
</tr>
<tr>
<td>Minnesota</td>
<td>8/1/74, 3/19/83</td>
<td>6, 12</td>
</tr>
<tr>
<td>Missouri</td>
<td>9/28/85</td>
<td>9</td>
</tr>
<tr>
<td>Nevada</td>
<td>1/1/88</td>
<td>18</td>
</tr>
<tr>
<td>New Hampshire</td>
<td>8/22/81</td>
<td>10</td>
</tr>
<tr>
<td>New Mexico</td>
<td>7/1/83</td>
<td>6</td>
</tr>
<tr>
<td>New York</td>
<td>1/1/86</td>
<td>6</td>
</tr>
<tr>
<td>North Carolina</td>
<td>1/1/82</td>
<td>3</td>
</tr>
<tr>
<td>North Dakota</td>
<td>7/1/83</td>
<td>10</td>
</tr>
<tr>
<td>Oklahoma</td>
<td>1/1/76</td>
<td>1</td>
</tr>
<tr>
<td>Oregon</td>
<td>1/1/82</td>
<td>6</td>
</tr>
<tr>
<td>Rhode Island</td>
<td>7/1/88</td>
<td>18</td>
</tr>
<tr>
<td>South Carolina</td>
<td>1/1/79, 1/1/90</td>
<td>2, 6</td>
</tr>
<tr>
<td>South Dakota</td>
<td>7/1/84, 3/3/88</td>
<td>3, 18</td>
</tr>
<tr>
<td>Tennessee</td>
<td>1/1/81</td>
<td>3</td>
</tr>
<tr>
<td>Texas</td>
<td>1/1/81</td>
<td>6</td>
</tr>
<tr>
<td>Utah</td>
<td>7/1/86</td>
<td>2</td>
</tr>
<tr>
<td>Vermont</td>
<td>5/14/86</td>
<td>6</td>
</tr>
<tr>
<td>Virginia</td>
<td>4/17/86</td>
<td>3</td>
</tr>
<tr>
<td>Wisconsin</td>
<td>5/14/80</td>
<td>18</td>
</tr>
</tbody>
</table>

**Sources:** Hewitt (1985), Thompson Publishing Group (1992), and state statutes.

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We consider only states with laws that cover both voluntary and involuntary separations, since we have only very noisy data on the nature of the separation.

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Details on state laws are from Hewitt (1985) and Thompson Publishing (1992) and have been cross-checked against the actual statutes. Table 1 lists only those states with laws that apply to employees who terminate their employment voluntarily.

Eighteen months is the maximum length of coverage available following the voluntary or involuntary termination of employment. COBRA also provides up to 36 months of coverage for family members who would otherwise lose their insurance coverage through events such as an employee's death, divorce from the employee, or the employee's eligibility for Medicare.
employees of religious organizations and the federal government were originally exempt from COBRA, although federal employees have subsequently been included (beginning in 1990). When the specific details of the state and federal statutes are at odds, firm provision of continuation benefits is governed by the law that provides for more generous coverage.

An important feature of continuation coverage from our perspective is the interaction between continuation coverage from an old job and insurance coverage from a new job. Before 1990, once individuals who were continuing their health insurance benefits from their old jobs obtained new employment that offered health insurance, the former employers could drop their continuation benefits, regardless of whether they were covered by the new policy. This is an important provision, because moving to a firm that offers insurance does not guarantee coverage if the firm excludes pre-existing conditions from its policy, or if it has a length of service requirement. After 1990, individuals were allowed to continue their COBRA coverage even if they started a new job with health insurance, as long as they continued to pay the required premiums. This change improved the portability features of COBRA, since it could now be used to bridge periods of non-coverage due to unemployment, length of service requirements, or pre-existing conditions exclusions.

The effective dates of the state laws are listed in Table 1. The federal coverage mandated under COBRA was phased in. Beginning in July 1986, firms had to offer continuation benefits at the start of their next plan year. For workers who were provided health insurance under union contracts, such benefits did not have to be offered until the next contract negotiation after January 1987.

Both the state and federal laws stipulate that the employee must pay the full cost of the coverage. At the federal level, full cost is defined specifically as 102% of the average employer cost of providing coverage. The coverage must be identical to that provided to similarly situated active employees, including the option to continue enrollment in supplemental insurance plans (such as for vision or dental care) if these are available. Although the continuation cost is substantially more than an individual generally had to pay as an employee, it may still be much more attractive than purchasing individual insurance, due to the economies of scale in administering group insurance and the reduced potential for adverse selection with large employee groups.

In Massachusetts, the average cost of family health insurance coverage per employee in 1989 was $3882. When inflated by the medical care component of the Consumer Price Index, this cost is equivalent to $5047 in 1993 dollars. In contrast, a New England commercial insurance company is offering a family policy for a 40-year-old man with a wife and two children, with a 1-year exclusion of pre-existing conditions, for $7000. Thus, an individual can realize non-trivial financial savings by purchasing continuation benefits rather than individual coverage.

Furthermore, for less healthy persons or those with high family medical expenses, individual coverage will be even less attractive, for two reasons. First, individual coverage is often medically underwritten, so that it may be much more expensive than the figure given above, or not available at all. Second, individual coverage is typically much less generous than group coverage. For example, the individual policy cited above excludes pre-existing conditions for one year, whereas such conditions would be covered from the start for someone continuing a group policy. Table 2 compares the health insurance benefits of individuals covered under group and nongroup policies in 1977. In every category, those covered under nongroup policies receive more

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8 Whether this was much of a change in practice is unclear, since it is not obvious how the old employer could verify that the employee had obtained new employment that offered health insurance coverage.

9 Authors' calculation using unpublished data from the Health Insurance Association of America.
Table 2. Group and Nongroup Health Insurance Benefits.

<table>
<thead>
<tr>
<th>Description of Benefits</th>
<th>Fraction of Individuals with Specified Benefits</th>
<th>Non-Group Plans</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Group Plans</td>
<td>Non-Group Plans</td>
</tr>
<tr>
<td>A. Primary Benefits</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Major medical coverage</td>
<td>86.9%</td>
<td>39.1%</td>
</tr>
<tr>
<td>Hospital room and board</td>
<td>98.4</td>
<td>91.4</td>
</tr>
<tr>
<td>Surgery</td>
<td>97.6</td>
<td>91.6</td>
</tr>
<tr>
<td>Physician office visit</td>
<td>87.9</td>
<td>40.4</td>
</tr>
<tr>
<td>B. Other Benefits</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Ambulance</td>
<td>89.0</td>
<td>54.0</td>
</tr>
<tr>
<td>Outpatient diagnostic services</td>
<td>95.9</td>
<td>66.0</td>
</tr>
<tr>
<td>Prescribed medicines</td>
<td>87.3</td>
<td>30.3</td>
</tr>
<tr>
<td>Mental health</td>
<td>92.2</td>
<td>66.0</td>
</tr>
<tr>
<td>C. Generosity of Benefits</td>
<td>(Conditional on Having Benefit)</td>
<td></td>
</tr>
<tr>
<td>Major medical deductible &lt; $100</td>
<td>94.3</td>
<td>61.6</td>
</tr>
<tr>
<td>Full semi-private room charge</td>
<td>77.8</td>
<td>38.2</td>
</tr>
<tr>
<td>80-100% of UCR surgical charge</td>
<td>70.6</td>
<td>60.0</td>
</tr>
<tr>
<td>80-100% of UCR physician charge</td>
<td>91.8</td>
<td>81.3</td>
</tr>
</tbody>
</table>

Source: Farley (1986), Tables 45–58.

Limited benefits. Relative to those with nongroup coverage, those with group policies are more than twice as likely to receive major medical coverage or coverage for physician office visits and prescription drugs, and more than 50% more likely to receive ambulance, mental health, and outpatient diagnostic service coverage. Furthermore, nongroup policies generally feature both higher deductibles and higher copayments.

Thus, whereas for a healthy person continuation coverage may provide only modest savings, for persons with high expected medical expenditures the difference in costs may be quite sizable. Since this population is exactly the one for which job-lock is potentially the largest problem, continuation mandates may have some effect in alleviating this distortion.

Table 3 shows the sources of health insurance coverage among 20–54-year-old men from the 1987 National Medical Expenditure Survey. Sixty-six percent of persons in this age group have employer-provided health insurance; most of that coverage is in their own name. Individuals who leave jobs are much less likely to have had insurance coverage on those jobs. This pattern could be construed as evidence of job-lock; on the other hand, it may just be that individuals are more likely to leave “bad” jobs that do not offer health insurance.

An important indication of the potential effectiveness of continuation mandates is the extent to which eligible persons avail themselves of these benefits. Using data from a company that administers COBRA claims, Flynn (1992) estimated that 19.3% of individuals who were terminated, were laid off, or quit from firms that had health insurance chose to take up continuation benefits. As Klirnan and Rahman (1992) noted, this estimate represents a lower bound on the take-up rate of benefits among the eligible population, since many of these persons will move to new jobs with health insurance or are covered by their spouses; these are individuals we would not expect to take up continuation benefits.

Data from the 1987 National Medical Expenditure Survey (NMES) suggest that about two-thirds of those who leave jobs with health insurance move to jobs in which they also receive health insurance. Klirnan and

10This figure excludes retirees. Take-up among retirees is likely to be substantially higher than that among individuals still working (see Gruber and Madrian 1993b).

11Authors' calculation. This figure actually represents an upper bound on the fraction who move to
Rahman estimate that an additional 10% of job leavers who do not find jobs with health insurance are covered by their spouse's health insurance.

Based on this information, we would expect approximately 30% of those who leave jobs with health insurance to be affected by COBRA. This figure implies COBRA take-up rates among those for whom we would expect it to matter of 65%. It therefore appears that a sizable fraction of eligible job changers avail themselves of continuation benefits, despite the high out-of-pocket costs.

Should we expect continuation mandates to alleviate job-lock? The answer depends on the nature of this distortion. If job-lock arises largely due to short-run medical considerations, such as pregnancy, then these policies may be quite effective. On the other hand, if individuals are worried about long-run coverage, then even 18 months of continuation benefits may not be sufficient. Furthermore, if job-lock arises from fear of moving to jobs with length of service requirements or pre-existing conditions exclusions, then temporary coverage may be all that is necessary, but if fear of being medically underwritten out of coverage is more influential, then more permanent portability is required. Finally, if job-lock primarily occurs among individuals with very high expected medical costs, then continuing to pay the cost of group health insurance will not be a major deterrent to the use of continuation benefits, but if job-lock represents risk aversion on the part of the average person, then continuation benefits may not seem a reasonable option financially. Thus, by examining the effects of continuation mandates on mobility, we hope to be able to draw some lessons about the nature of the job-lock problem itself.

Data and Regression Framework

The data we use come from the 1984, 1985, 1986, and 1987 panels of the Survey of Income and Program Participation (SIPP). This nationally representative survey collected information on the economic and demographic characteristics of individuals and their families through a series of quarterly interviews, referred to as "waves," for roughly 2-1/2 years. Because the panels overlap, as many as three panels may be interviewed concurrently. Altogether, these four panels span the period from June 1983 to April 1989.

During each interview, respondents were asked questions concerning up to two jobs held during the previous four calendar months, including questions about the industry, occupation, hours worked, and pay of each job, as well as the start or end date for each job if it was held for less than the full quarter. With this information, we are able determine whether an individual changed jobs or ceased to be employed during the quarter.

Our sample is restricted to men between the ages of 20 and 54 who were not in school over the course of their participation in the SIPP and who were not self-employed. We exclude individuals from jobs with health insurance, as the NMES does not distinguish between those who receive health insurance through their current employer and those who receive COBRA benefits through a former employer. Prior evidence on this question is mixed. Madrian (1994) found job-lock arising from both pregnancy, which is a short-run expenditure, and larger families, which gives rise to longer-run expenditures. Holtz-Eakin (1994) found statistically significant evidence of job-lock only over short periods of observation, however; job changes over a period of more than one year were unaffected by the provision of health insurance.

Note, however, that in this case continuation benefits in the period after 1990 would be much more effective.

14The 1984 Panel of the SIPP, which has nine waves, covers three years. The 1985 Panel consists of eight waves (32 months), and the 1986 and 1987 Panels both include seven waves.
15Although we would like to be able to study the 14 state continuation of coverage laws that were passed before 1983, the period covered by the data provides a window around the passage of the federal law and includes 19 state laws that took effect after June 1983.
16We exclude 55-64-year-olds because the effects of continuation mandates on retirement behavior may be quite different from their effects on job mobility. We focus on the older group of workers in Gruber and Madrian (1993a, 1993b), finding sizable effects of continuation mandates on retirement. We exclude women because the process determining their job mobility may be quite different from that for men.
Hawaii, which has mandated health insurance for all employees, and West Virginia, for which we were unable to definitively date the continuation mandate. Individuals from several other small states are also excluded because, out of concern for confidentiality, the SIPP has grouped these states together, making it impossible to assign the appropriate state laws to individuals in these states.\textsuperscript{17} The final sample consists of 155,151 quarterly observations on 29,841 individuals.

We use these data to estimate a model of quarterly job turnover. Turnover is defined as changing employers, becoming self-employed, or becoming unemployed.\textsuperscript{18} Overall, the quarterly turnover rate for our sample is 10%, and 24% of each panel changes jobs within the first year. This figure is similar to the annual turnover rate of 22.6% that is estimated from the 1987 National Medical Expenditure Survey. It should be noted that we are not able to separate voluntary from involuntary job changes, because the 1984 and 1985 panels of the SIPP did not ask about the reason for leaving one's job. Although job-lock really applies only to voluntary turnover, we do not think that this problem seriously biases our results. Over the time period that we consider, late 1983 to early 1989, the economy was growing and the majority of turnover taking place was voluntary.\textsuperscript{19} Furthermore, Madrian (1994) found that her estimates of job-lock were not sensitive to whether or not her job change variable included those who changed jobs involuntarily.

To calculate the wages associated with the job that an individual left, we divide the quarterly wage and salary earnings of the job by the usual hours worked per week multiplied by the number of weeks in the quarter for which an individual was employed in that job. The wage rate computed in this manner is highly correlated with the hourly wage rate reported by those who are paid hourly.

Each individual is assigned the maximum months of continuation coverage available at the beginning of the wave under either COBRA or the appropriate state law. Because the federal law was phased in between July 1986 and June 1987, we increase the months of coverage available under COBRA by 1.5 months for each month over this period. By June 1987, therefore, months of continuation coverage equals 18 for all individuals except those who live in Connecticut, where state law allows for 20 months of coverage.

Because continuation coverage is only available for those with employer-provided health insurance, an important variable in our analysis is whether or not an individual is covered by such insurance. In each wave, the SIPP asks if an individual is covered by health insurance, if this coverage comes from a policy in one's own name, and, if so, if this policy is provided by an employer or union. In quarters in which individuals do not change jobs, we code them as having employer-provided health insurance if they report being covered by health insurance provided by an employer or union. In quarters in which individuals change jobs and report coverage from employer-provided health insurance, it is not clear from the SIPP questions whether the employer-provided health insurance was attached to the first job, the second job, or both. We code these individuals' coverage by employer-provided health insurance based on whether or not they report having such coverage in the previous wave of the SIPP.\textsuperscript{20}

The means for our data sample are pre-

\textsuperscript{17}These states are Alaska, Idaho, Iowa, Maine, Mississippi, Montana, New Mexico, North Dakota, South Dakota, Vermont, and Wyoming.

\textsuperscript{18}Although our definition of turnover includes the switch from employment to self-employment, it does not include moving from one self-employed job to another, since individuals are excluded from the sample when they are self-employed.

\textsuperscript{19}Three-quarters of job changes in the 1987 National Medical Expenditure Survey were voluntary (see Madrian 1994).

\textsuperscript{20}This definition of coverage by employer-provided health insurance will still present problems in the case of individuals who change jobs in two consecutive quarters. Of the job changes in our sample, 15% are immediately preceded by another quarter with a job change, so only 1.5% of our sample are
sent in Table 4. Each observation is a quarter in which a member of our sample does any work. \textsuperscript{21} Eleven percent of the sample is non-white, and the average educational level is approximately one year of college. Seventy-seven percent of the sample holds employer-provided health insurance. As noted above, the quarterly turnover rate is 10%.

The model that underlies the analysis is presented in detail in Madrian (1994). Different individuals value health insurance at different levels, which leads to job-lock through the mechanism described above. We assume that these valuations are distributed randomly in the population we are studying.\textsuperscript{22} We can examine the effectiveness of continuation benefits in alleviating job-lock through a probit regression of the form

\begin{equation}
\Pr(\text{Leave Job}) = \Phi(\alpha + \beta_1 \cdot X_{ij} + \beta_2 \cdot State_j + \beta_3 \cdot Time + \beta_4 \cdot Law_j),
\end{equation}

where \(i\) indexes individuals, \(j\) indexes states, and \(t\) indexes time. \(X_{ij}\) is a set of individual demographic and job characteristics, \(State\) is a set of state dummies, \(Time\) is a set of year and month dummies,\textsuperscript{23} and \(Law_j\) is the number of months of continuation coverage available in state \(j\) at time \(t\). The regression models the probability that an individual leaves his job in a particular quarter.

We control for a number of characteristics of individuals (education, experience, and race) and their jobs (log hourly wage, eight industry dummies, six occupation dummies, and a dummy for the provision of health insurance). The year dummies control for any national time trends in mobility, and the month dummies control for seasonal mobility patterns. The state dummies control for any time-invariant differences across states that may be correlated with their propensity to legislate continuation mandates, such as differences in the underlying valuation of health insurance among the population. The regressor of interest measures the number of months of continuation coverage for which workers are eligible if they leave their current jobs. If continuation of coverage mandates alleviate job-lock, this regressor should have a positive coefficient.

### Effects on Mobility

#### Basic Results

The basic mobility results are presented in Table 5. The covariates have their expected effects. Well-educated and older workers are less likely than other workers to

---

**Table 4. Descriptive Statistics.**

<table>
<thead>
<tr>
<th>Variable</th>
<th>Mean</th>
<th>Standard Deviation</th>
</tr>
</thead>
<tbody>
<tr>
<td>Non-white</td>
<td>0.113</td>
<td>0.316</td>
</tr>
<tr>
<td>Years of School Completed</td>
<td>12.94</td>
<td>3.021</td>
</tr>
<tr>
<td>Experience</td>
<td>16.42</td>
<td>10.044</td>
</tr>
<tr>
<td>Log Hourly Wage</td>
<td>$2.20</td>
<td>0.619</td>
</tr>
<tr>
<td>Holds Employer-Provided Health Insurance</td>
<td>0.772</td>
<td>0.420</td>
</tr>
<tr>
<td>Leave Job</td>
<td>0.101</td>
<td>0.301</td>
</tr>
<tr>
<td>Months of Continuation Coverage</td>
<td>6.88</td>
<td>7.257</td>
</tr>
</tbody>
</table>

**Authors' calculations using data on men aged 20-54 from the 1984–87 panels of the Survey of Income and Program Participation. The sample consists of 155,151 quarterly observations on 29,841 individuals.**

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\textsuperscript{21}Thus, there are multiple observations on each individual. Accounting for the resultant intra-personal correlation in the error term has little effect on the results. There is also a potential left-censoring problem that may lead to the type of "dynamic sample-selection" bias noted by Diamond and Hausman (1984). Individuals who left their jobs in response to the availability of continuation benefits, but who were still unemployed one quarter later, will be excluded from our sample. In a bivariate setting, this exclusion would impart a downward bias to our estimated effect; in a multivariate setting, the effect is unclear. We tested for the importance of this bias in our study of the effect of continuation mandates on retirement (Gruber and Madrian 1993a) and found it to be small.

\textsuperscript{22}Ideally, one would account for these individual differences in valuation and use them to help identify any potential effects of continuation benefits. We discuss one attempt to do so below.

\textsuperscript{23}The month dummies are actually dummies for the month in which the quarter begins. The year dummies are dummies for the year in which the panel begins.
Table 5. Effect of Continuation Coverage on Job Turnover.

<table>
<thead>
<tr>
<th>Independent Variable</th>
<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Non-White</td>
<td>0.0118</td>
<td>-0.0087</td>
<td>-0.0087</td>
</tr>
<tr>
<td></td>
<td>(0.0142)</td>
<td>(0.0144)</td>
<td>(0.0144)</td>
</tr>
<tr>
<td>Education</td>
<td>-0.0130</td>
<td>-0.0072</td>
<td>-0.0071</td>
</tr>
<tr>
<td></td>
<td>(0.0019)</td>
<td>(0.0020)</td>
<td>(0.0020)</td>
</tr>
<tr>
<td>Experience</td>
<td>-0.0127</td>
<td>-0.0106</td>
<td>-0.0106</td>
</tr>
<tr>
<td></td>
<td>(0.0005)</td>
<td>(0.0005)</td>
<td>(0.0005)</td>
</tr>
<tr>
<td>Log Hourly Wage</td>
<td>-0.2645</td>
<td>-0.1527</td>
<td>-0.1338</td>
</tr>
<tr>
<td></td>
<td>(0.0084)</td>
<td>(0.0084)</td>
<td>(0.0084)</td>
</tr>
<tr>
<td>Health Insurance</td>
<td>-0.5546</td>
<td>-0.5542</td>
<td>0.0046</td>
</tr>
<tr>
<td></td>
<td>(0.0108)</td>
<td>(0.0108)</td>
<td>(0.0011)</td>
</tr>
<tr>
<td>Months of Coverage</td>
<td>0.0046</td>
<td>0.0046</td>
<td>0.0046</td>
</tr>
<tr>
<td></td>
<td>(0.0011)</td>
<td>(0.0011)</td>
<td>(0.0011)</td>
</tr>
<tr>
<td>Log-Likelihood</td>
<td>-47,098</td>
<td>-45,794</td>
<td>-45,785</td>
</tr>
<tr>
<td></td>
<td>(155,151)</td>
<td>(155,151)</td>
<td>(155,151)</td>
</tr>
</tbody>
</table>

B. Marginal Effect of 12 Months of Continuation Coverage [% of Baseline Mobility Rate]

|                      | 0.0087    | [9.1%]    |
|                      | (0.0011)  |           |

Notes: The table gives estimates from a probit equation for whether or not an individual changed jobs in a quarter using data from the 1984–87 Survey of Income and Program Participation. The sample is comprised of men aged 20–54. Coefficients for panel, state, month, industry, and occupation dummies are not reported.

The coefficient on the months of continuation coverage variable in the third column is highly significant. In this table and the remaining tables, the probability derivative for this coefficient is presented in the last row. It indicates that one year of continuation benefits increases the mobility rate for individuals in our sample by 0.9 percentage points. This effect represents approximately 9% of the baseline quarter-to-quarter mobility rate for the sample. Although small in absolute magnitude, this effect is quite sizable relative to Madrian’s (1994) 25% estimate of job-lock; it implies that a year of continuation coverage reduces job-lock by almost 40%.

Specification Checks

The key identifying assumption behind our model is that the passage of these mandates represents an exogenous change in the opportunity set of the worker making a mobility decision. An alternative hypothesis is that the laws themselves are endogenous responses to the pattern of mobility across states. For example, if more individuals are changing (or losing) jobs, states may respond by mandating benefits that cover individuals during this transition. To the extent that the state mandates are responses to long-standing mobility differences across states, the state fixed effects included in the regression will control for this endogeneity. To the extent that legislatures are responding to changes in mobility, however, state fixed effects will not be sufficient controls.

25Madrian (1994) also found that the wage coefficient fell substantially when health insurance was added as a regressor. Mitchell (1982) found a similar result for pensions.

24 We should note that Madrian (1994) produced three estimates of job-lock corresponding to three different populations: individuals who had a spouse with health insurance, those with families of a certain size, and those who had a pregnant spouse. The 25% estimate of job-lock cited here is the estimate for individuals who had a spouse with health insurance. This is the right estimate to consider in the context of this paper, because it is based on only the availability of health insurance, rather than (as the other two estimates are) expected medical expenditures. Continuation coverage increases the availability of health insurance to all individuals regardless of their family size or other factors affecting the need for health insurance.
There is a natural test for this alternative hypothesis. Because only those workers who have health insurance on the job are eligible to receive continuation benefits, workers without health insurance on the job provide a control group for assessing the effects of continuation mandates. If these mandates are simply correlated with, or due to, exogenous changes in mobility propensities, then the laws should be correlated with the mobility of workers both with and without health insurance on the job. If the laws are causing changes in mobility patterns, however, they should only affect workers with employer-provided health insurance.

Table 6 therefore divides the sample into two groups consisting of those with and those without employer-provided health insurance on their current jobs. For those with health insurance (column 1), the effect of continuation coverage is even greater than that estimated for the full sample, and it is more significant as well. The coefficient estimate implies that one year of continuation coverage leads to an increase in the likelihood of changing jobs of .84 percentage points; this is 14% of the baseline mobility rate for this group.

The second column of Table 6 looks at those without health insurance. In fact, the mandates positively affect mobility for this group as well. This finding suggests that the passage of the continuation regulations was correlated with other factors causing increases in mobility in these states during the years under study. The implied mobility increase for those without health insurance, however, is not significant, and it represents only 3.4% of the baseline mobility rate for this group. Furthermore, if we use the estimated effect of these mandates on mobility for those without health insurance to control for secular changes in mobility, the evidence in Table 6, consistent with that in Table 5, suggests that continuation mandates increase mobility by about 10% (that is, 14% minus 3.4%).

As noted earlier, the effect of continuation mandates will be mitigated by factors such as the extent to which firms self-insure (since self-insured firms are exempt from state mandates) and the distribution of workers by firm size (since small firms are exempt from the federal mandate). In order to appropriately measure the effect of continuation coverage, we need to account for variation in these factors across states. We do so by constructing "corrected" months of coverage variables. The Appendix details the calculation of these correction factors, one for those who have employer-provided health insurance that adjusts for the factors just described, and one for all individuals that also adjusts for the likelihood of being insured.

Briefly, we use unpublished data from the Health Insurance Association of America (HIAA) and the Current Population Survey to estimate, by state, the fraction of workers actually offered employer-provided health insurance, the fraction in self-insured firms, and the fraction in small (\(<\ 20\) employee) firms. If only a state law is in effect, we adjust the months of coverage

Table 6. Effect of Continuation Coverage on Job Turnover.

<table>
<thead>
<tr>
<th>Independent Variable</th>
<th>Have Health Insurance</th>
<th>No Health Insurance</th>
</tr>
</thead>
<tbody>
<tr>
<td>A. Coefficient Estimates</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Non-White</td>
<td>.0038</td>
<td>-.0238</td>
</tr>
<tr>
<td></td>
<td>(.0195)</td>
<td>(.0216)</td>
</tr>
<tr>
<td>Education</td>
<td>-.0105</td>
<td>-.0026</td>
</tr>
<tr>
<td></td>
<td>(.0026)</td>
<td>(.0030)</td>
</tr>
<tr>
<td>Experience</td>
<td>-.0103</td>
<td>-.0103</td>
</tr>
<tr>
<td></td>
<td>(.0007)</td>
<td>(.0008)</td>
</tr>
<tr>
<td>Log Hourly Wage</td>
<td>-.1770</td>
<td>-.0862</td>
</tr>
<tr>
<td></td>
<td>(.0118)</td>
<td>(.0121)</td>
</tr>
<tr>
<td>Months of Coverage</td>
<td>.0059</td>
<td>.0021</td>
</tr>
<tr>
<td></td>
<td>(.0014)</td>
<td>(.0018)</td>
</tr>
<tr>
<td>Log-Likelihood</td>
<td>-27,180</td>
<td>-18,469</td>
</tr>
<tr>
<td>N</td>
<td>119,700</td>
<td>35,451</td>
</tr>
<tr>
<td>B. Marginal Effect of 12</td>
<td>.0084</td>
<td>.0075</td>
</tr>
<tr>
<td>Months of Continuation</td>
<td>[14.3%]</td>
<td>[3.4%]</td>
</tr>
<tr>
<td>Coverage [% of Baseline Mobility Rate]</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Notes: See notes to Table 5.

\(^{26}\)Note that this split is imperfect due to potential errors in assigning insurance coverage to a given job for persons who change jobs in consecutive quarters.
Table 7. Effect of Continuation Coverage on Job Turnover (Corrected).

<table>
<thead>
<tr>
<th>Independent Variable</th>
<th>Full Sample</th>
<th>Have Health Insurance</th>
</tr>
</thead>
<tbody>
<tr>
<td>A. Coefficient Estimates</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Non-White</td>
<td>-.0088</td>
<td>.0038</td>
</tr>
<tr>
<td></td>
<td>(.0144)</td>
<td>(.0195)</td>
</tr>
<tr>
<td>Education</td>
<td>-.0072</td>
<td>-.0105</td>
</tr>
<tr>
<td></td>
<td>(.0020)</td>
<td>(.0026)</td>
</tr>
<tr>
<td>Experience</td>
<td>-.0106</td>
<td>-.0103</td>
</tr>
<tr>
<td></td>
<td>(.0005)</td>
<td>(.0007)</td>
</tr>
<tr>
<td>Log Hourly Wage</td>
<td>-.1339</td>
<td>-.1769</td>
</tr>
<tr>
<td></td>
<td>(.0084)</td>
<td>(.0118)</td>
</tr>
<tr>
<td>Health Insurance</td>
<td>-.5542</td>
<td>—</td>
</tr>
<tr>
<td></td>
<td>(.0108)</td>
<td></td>
</tr>
<tr>
<td>Months of Coverage</td>
<td>.0063</td>
<td>.0063</td>
</tr>
<tr>
<td></td>
<td>(.0016)</td>
<td>(.0017)</td>
</tr>
<tr>
<td>Log-Likelihood</td>
<td>-45,785</td>
<td>-27,181</td>
</tr>
<tr>
<td>N=Sample Size</td>
<td>155,151</td>
<td>119,700</td>
</tr>
<tr>
<td>B. Marginal Effect of 12</td>
<td>.0122</td>
<td>.0091</td>
</tr>
<tr>
<td>Months of Continuation</td>
<td>[12.7%]</td>
<td>[15.3%]</td>
</tr>
<tr>
<td>Coverage [% of Baseline</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Mobility Rate</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Notes: See notes to Table 5.

to reflect the fact that the state law applies only to those offered insurance and only to those who do not work in a firm that self-insures. If only a federal law is in effect, we adjust the months of coverage by the fraction of individuals who are offered insurance and the fraction who work in small firms. When both a state law and a federal law are in effect, the corrected months of coverage account for the fact that the state law will affect insured individuals working in small firms that do not self-insure. When we adjust for these factors, the average length of continuation coverage in states that offer coverage falls from 6.88 months to 4.55 months.

The results using the corrected months of coverage variables are presented in Table 7. The findings are very similar; the magnitudes of the coefficients are somewhat increased, as are the standard errors, although the effects are still highly significant. This result is to be expected from a procedure that multiplies our coefficient of interest by a noisy correction factor that is less than one. The estimates in columns 1 and 2 of Table 7 suggest that one year of coverage increases mobility by between 12% and 15%.27

Finally, one potential problem with this model that was discussed above is that the variable of interest is measured with considerable noise. In particular, we are not focusing on the individuals who most value their health insurance benefits and for whom the limited portability of continuation mandates may be the most useful. One such group for whom these mandates should be important is married workers, whose higher family medical expenditures may increase their valuation of health insurance. It is true that married workers are more likely than single workers to have coverage from an alternative source (namely their spouse), and thus potentially less likely to need continuation benefits. Madrian (1994) estimated that 33% of married men had coverage through a spouse’s employment in 1987. But family health care expenditures for married men are three times greater than those for single men.28 Thus, on net, being married would appear to raise the costs of leaving a job with health insurance.29

In results not reported, we repeated the regressions of Tables 5–7 for only married individuals and found slightly larger, although less precisely estimated, effects. For the basic specification, there is a .0085 percentage point increase in mobility, which is 9.7% of the baseline mobility rate for this subsample. This result implies that the mandates had a slightly larger effect on a

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27This regression is not run for those without health insurance, since their corrected months of continuation coverage are zero by definition.
28Authors’ tabulations from the 1980 National Medical Care Utilization and Expenditure Survey.
29Of course, if these higher expected expenditures are reflected in lower wages, the mobility distortion for married men will be no higher than that for single men. The feasibility of wage shifting by marital status is uncertain; see Gruber (1994) for evidence in favor of the ability of employers to shift increased insurance costs to the wages of identifiable demographic groups within the workplace.
population that was more likely to be affected by their presence.

Continuation Mandates and Insurance Coverage

A necessary (but not sufficient) precondition for continuation of coverage mandates to affect job mobility is that they affect insurance coverage. Moreover, the effects of these mandates on insurance coverage may be important for assessing their welfare implications. Ostensibly, the purpose of these mandates was to correct a failure in the market for private insurance for job leavers. Such market failure seems likely, given the adverse selection in the take-up of continuation coverage benefits documented by Huth (1991) and Long and Marquis (1992). In this case, increasing insurance coverage among those who would have left their jobs even in the absence of the regulations represents a clear welfare improvement. As discussed earlier, however, the welfare implications of reducing job-lock are less clear. Thus, the net welfare effects of the mandates may depend on the relation between their effect on “inframarginal” individuals, who would have left their jobs in the absence of a mandate, and their effects on “marginal” individuals whose mobility decision is made in response to the mandate.

To model the effect of these mandates on insurance coverage, we examine the sample of workerquarters in which there was a job change. We then create a dummy variable equal to one if an individual has insurance coverage in the quarter after his job change. We regress this dummy variable on the set of covariates used earlier.30

Table 8. The Effect of Continuation Mandates on Insurance Coverage.

<table>
<thead>
<tr>
<th>Independent Variable</th>
<th>Full Sample</th>
<th>Have Health Insurance</th>
<th>No Health Insurance</th>
</tr>
</thead>
<tbody>
<tr>
<td>A. Coefficient Estimates</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Non-White</td>
<td>-.2961</td>
<td>-.3227</td>
<td>-.1392</td>
</tr>
<tr>
<td>(0.0350)</td>
<td>(0.0617)</td>
<td>(0.0500)</td>
<td></td>
</tr>
<tr>
<td>Education</td>
<td>.1242</td>
<td>.0929</td>
<td>.1179</td>
</tr>
<tr>
<td>(0.0046)</td>
<td>(0.0077)</td>
<td>(0.0070)</td>
<td></td>
</tr>
<tr>
<td>Age</td>
<td>.0112</td>
<td>.0368</td>
<td>-.0289</td>
</tr>
<tr>
<td>(0.0096)</td>
<td>(0.0174)</td>
<td>(0.0146)</td>
<td></td>
</tr>
<tr>
<td>Age²</td>
<td>.0001</td>
<td>-.0002</td>
<td>.0005</td>
</tr>
<tr>
<td>(0.0001)</td>
<td>(0.0002)</td>
<td>(0.0002)</td>
<td></td>
</tr>
<tr>
<td>Months of Coverage</td>
<td>.0034</td>
<td>.0112</td>
<td>-.0048</td>
</tr>
<tr>
<td>(0.0027)</td>
<td>(0.0048)</td>
<td>(0.0043)</td>
<td></td>
</tr>
<tr>
<td>Log-Likelihood</td>
<td>-7,535</td>
<td>-2,513</td>
<td>-3,347</td>
</tr>
<tr>
<td>N</td>
<td>12,285</td>
<td>4,959</td>
<td>5,070</td>
</tr>
<tr>
<td>B. Marginal Effect of 12 Months of Continuation Coverage</td>
<td>.0145</td>
<td>.0411</td>
<td>-.0219</td>
</tr>
</tbody>
</table>

Notes: This table gives estimates of the probability of being insured in the quarter after a job change using data from the 1984-87 Survey of Income and Program Participation. The sample is comprised of men aged 20-54 who change jobs or become unemployed. Coefficients for panel dummies are not reported.

These results are presented in Table 8. Since there are relatively few individuals in some states, we run the regressions without state effects; the coefficients are quite similar, although less precisely estimated, if state effects are included. Once again, we use a probit specification.

For the full sample (column 1), the coefficient on months of coverage suggests that continuation mandates have a positive but insignificant effect on the probability of being insured after changing jobs. The implied increase in insurance coverage is 1.4 percentage points. In the second and third columns of Table 8, we split the sample, as before, into those with and without health insurance. For those with health insurance, the effect of continuation coverage is statistically significant and implies a 4.1 percentage point increase in insurance coverage after changing jobs. For those without employer-provided health insur-

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30 Some evidence that these mandates did increase coverage among those who left their jobs and remained unemployed is presented in Klerman and Rahman (1992). For comparability to our mobility regressions, however, we wish to examine the effects on the insurance coverage for any worker who leaves his job.

31 We do not include the job characteristics (hourly wage, industry, and occupation), since it is not obvious which job (the one left or the one gained) to use. The results are similar if the characteristics of the former job are included.
ance, the likelihood of having insurance after leaving a job falls by 2.2 percentage points. Thus, the effect for those with health insurance is strengthened if we consider other trends correlated with the passage of the mandates. Compared to the effect on mobility for those with health insurance (0.84 percentage points), this result suggests that the effect of the mandates was largely to cover individuals who would have left their jobs even in the absence of continuation availability.

Conclusion

The problem of job-lock has caught the public's attention, and one key measure of the success of health care reform in the United States will be its ability to mitigate insurance-induced mobility reductions. That goal, however, is only one of several toward which reform must strive, goals that may not be mutually compatible. If it is possible to substantially alleviate job-lock through modest policy changes, doing so may offer policy-makers more degrees of freedom with which to design other elements of the reform package.

This research suggests that continuation benefits have had some success in alleviating job-lock. We find that one year of continuation benefits was associated with a 10% increase in mobility among those with health insurance—a moderately large effect if, as Madrian (1994) has estimated, insurance-induced job-lock reduces mobility by about 25%. Furthermore, above and beyond increasing job mobility, continuation mandates were successful in increasing the insurance coverage of job leavers.

Our findings therefore imply that a substantial portion of the job-lock problem can be alleviated by mandatory limited portability. This conclusion is supportive of Holtz-Eakin's (1994) contention that job-lock is a short-run problem. Presumably, the expansion of eligibility for COBRA coverage after 1990, allowing continued health coverage through the previous employer even after an individual finds employment with a new firm that offers health insurance (for which there may be length of service requirements or pre-existing conditions exclusions), will strengthen the effects of COBRA. Future work to model the effects of this change would be very useful.

Although our estimates imply that limited portability significantly reduces job-lock, there are a number of ways in which limited portability may need to be refined to mitigate this distortion for the entire population. The important point from an economic perspective is that the policy changes required to fully remove job-lock may be very costly relative to the changes necessary to greatly mitigate its effects. If limited portability has a substantial effect, policies that strengthen COBRA, while perhaps providing for a public reinsurance pool for the worst risks, may be all that are necessary to greatly reduce this key labor market distortion.

APPENDIX

Calculating Correction Factors for the Impact of State and Federal Continuation Laws

In our basic regression specification, we assign to individuals the maximum number of months of continuation coverage mandated under either federal or state law. Several factors, however, lead to less than full coverage of these laws. First, those who are not covered by employer-provided health insurance should not be affected by any of these laws. Second, before the federal law took effect, employees of self-insured firms should not have been affected by the state laws because, under the 1974 Employee Retirement Income and Security Act (ERISA), self-insured firms are exempt from state mandates. And third, those who work in firms with fewer than 20 employees will not be affected by the federal law, which does not apply to small firms, although they will be influenced by state laws that pertain to all firms.

To the extent that these factors differ across states, we would expect laws that mandate equivalent months of coverage to have different effects. To account for the less than full coverage of these laws, we compute a "corrected" measure of months of coverage. As before, if there is no state or federal law in effect, we
assign no months of coverage to an individual. If a state law is operative but the federal law has not yet taken effect, the corrected months of coverage equals

$\left( \text{Months of State Coverage} \right) \cdot \left[ \left( \frac{\text{Fraction Offered Insurance}}{} \right) \cdot \left( \frac{\text{Fraction in Self-insured Firm}}{1-\frac{\text{Fraction Insured in <20 Firm}}{\text{Fraction Insured in <20 Firm}}} \right) \right]$. 

If the federal law is in effect but there is no state law, the corrected months of coverage equals

$\left( \text{Months of Federal Coverage} \right) \cdot \left[ \left( \frac{\text{Fraction Offered Insurance}}{\text{Self-insured}} \right) \cdot \left( \frac{\text{Fraction Insured in <20 Firm}}{1-\frac{\text{Fraction Insured in <20 Firm}}{\text{Fraction Insured in <20 Firm}}} \right) \right]$. 

Finally, if both a state law and the federal law are in place, the corrected months of coverage equals

$\left( \text{Months of Federal Coverage} \right) \cdot \left[ \left( \frac{\text{Fraction Offered Insurance}}{\text{Self-insured}} \right) \cdot \left( \frac{\text{Fraction in Self-insured Firm}}{1-\frac{\text{Fraction Insured in <20 Firm}}{\text{Fraction Insured in <20 Firm}}} \right) \right] + \left( \text{Months of State Coverage} \right) \cdot \left[ \left( \frac{\text{Fraction Offered Insurance}}{\text{Self-insured}} \right) \cdot \left( \frac{\text{Fraction in Self-insured Firm}}{1-\frac{\text{Fraction Insured in <20 Firm}}{\text{Fraction Insured in <20 Firm}}} \right) \right]$. 

We measure the various components of the corrected months of coverage using data from the Health Insurance Association of America (HIAA) 1989 employer survey and the May 1988 Current Population Survey (CPS) pension supplement. In using data from this late date, we are assuming that these factors are constant over time. This assumption clearly is incorrect for self-insurance, which grew dramatically during the 1980s. As long as the pattern of growth of self-insurance was uncorrelated with the passage of these laws, however, it will not bias our results. Recent research suggests that mandates play little role in the firm’s decision to self-insure.

The fraction of workers offered employer-provided health insurance and the fraction of insured workers in firms with fewer than 20 employees are measured directly, by state, from the CPS. The fraction in self-insured firms is estimated from the HIAA data. In using the HIAA data, we have tried two different approaches. The first is to calculate the average of these quantities by state. This strategy, however, is hindered by very small state cell sizes in the HIAA data. Our second approach, therefore, is to use the HIAA data to run a regression predicting the rate of self-insurance as a function of firm size, industry, and census division. We then use these estimated coefficients to impute values of these quantities for each individual in the May CPS. Our correction factors are state-wide averages of these imputed values.

We imputed values in this way for all male workers between 20 and 54 years old. In our empirical work, we use the corrected months of coverage computed as outlined above when the sample includes all workers. When the sample is restricted to only those with health insurance, we use a similarly constructed corrected months of coverage that does not adjust for the fraction of workers in the state with employer-provided health insurance.

REFERENCES


